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# UNEMPLOYMENT, HYSTERESIS AND TRANSITION

*Miguel A. León-Ledesma* and Peter McAdam***

## Abstract

*In this paper, we quantify the degree of persistence in the unemployment rates of transition countries using a variety of methods benchmarked against the EU. Initially, we work with the concept of linear 'Hysteresis' as described by the presence of unit roots in unemployment as in most empirical research on this area. Given that this is potentially a narrow definition, we also take into account the existence of structural breaks and nonlinear dynamics in unemployment. Finally, we examine whether CEECs' unemployment presents features of multiple equilibria, that is, if it remains locked into a new level whenever some structural change or sufficiently large shock occurs. Our findings show that, in general, we can reject the unit-root hypothesis after controlling for structural changes and business-cycle effects, but we can observe the presence of a high and low unemployment equilibria. The speed of adjustment is faster for CEECs than the EU, although CEECs tend to move more frequently between equilibria.*

## I Introduction

One of the foremost features of the transition process of Central and Eastern European Countries (CEECs) is the appearance of open unemployment hidden during the central planning regime. As reported in EBRD (2000) and IMF (2000), this phenomenon has had a deep impact on poverty and social exclusion. This is partly because the comprehensive social safety net left agents with little experience in dealing with the uncertainty and adversity associated with protracted unemployment. The labour-market reforms of the first half of the 1990s, especially the reduction of unemployment benefits, did not seem to have the expected impact in reducing unemployment by improving matching (Boeri, 1997a).[1] Employment expanded at a much slower pace than output, pointing to a high degree of persistence in unemployment, thus aggravating the social problems associated with the transition to a market economy. Furthermore,

*University of Kent
**European Central Bank

[1] For reviews of labour-market developments in Transition Economies see also EBRD (2000), Nesporova (2002) and Vidovic (2001), and Tichit (2000) for a comparative study of unemployment dynamics among Eastern European countries.

with the prospect of EU membership, accession countries will continue to pursue both product and labour-market reforms that are likely to exert important shocks on employment (EBRD, 2000). This is especially true if labour hoarding is reduced by the introduction of greater competition. Shocks are also likely to come about for some countries because of macroeconomic stabilization measures (i.e. budgetary consolidation, inflation and exchange rate stabilization) to meet EU membership requirements.

In this paper, we quantify the degree of persistence in the unemployment rates of transition countries using a variety of methods. As far as we are aware, this is the first systematic attempt to do so. Initially, we work with the concept of linear 'Hysteresis' as described by the presence of unit roots in unemployment as in most empirical research on this area. Given that this is potentially a narrow definition,[2] we also take into account the existence of structural breaks and nonlinear dynamics in unemployment. Finally, we examine whether CEECs'[3] unemployment presents features of multiple equilibria: that is, if it remains locked into a new level of unemployment whenever a structural change or sufficiently large shock occurs.

The question addressed is important for several reasons. First, it has implications for social protection and labour market reforms, as well as macro-stabilization policy in the CEECs. The presence of hysteretic or highly persistent unemployment would imply that unemployment could become a long-lasting problem after radical reforms. For instance, for countries showing features of multiple equilibria, reforms aimed at reducing non-employment benefits could constitute large and long-lasting positive shocks if done during an employment recovery phase. However, reforms carried out during rising unemployment may not have the desired effect of changing equilibrium unemployment as the positive labour market reform shock could be choked off by the negative (demand or supply) shock affecting the economy. Second, it helps to understand if the behaviour of unemployment in our set of countries is consistent with recently developed models of labour markets in transition. Comparison with the persistence profile in the EU could also help analyze the possible impact of common shocks. For instance, if unemployment were to be more persistent in the CEECs than in the EU, common negative shocks could increase migration pressures westwards, and common positive shocks reduce them.

To undertake our analysis, we first work with the concept of Hysteresis as stemming from the presence of a unit or near-unit root in unemployment rates. We apply a battery of unit root tests on a set of 12 CEECs (benchmarked against an EU-15 aggregate) to test for the existence of random-walk behaviour, quantify the degree of persistence and account for possible breaks in our sample and lack of power in our tests. Secondly, we use Markov-switching regressions to analyze persistence, taking into account the possibility of a changing

[2] For discussions of the concept and implications of Hysteresis in unemployment see Amable *et al.* (1995) and Cross (1995).

[3] CEECs in our sample comprise Poland, Romania, Slovenia, Croatia, Hungary, Bulgaria, Czech Republic, Slovakia, Estonia, Latvia, Lithuania and Russia.

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equilibrium unemployment due to breaks and business-cycle fluctuations. This, most importantly, allows us to work with a concept of Hysteresis as multiple equilibria in unemployment. In the next section, we provide an overview of the evolution of unemployment in our sample of CEECs and some theoretical models attempting to explain it.

## II Stylized Facts and Theoretical Background

The evolution of unemployment in Eastern Europe has showed a diversity of patterns (Figure 1). Overall, however, we can observe high levels of unemployment during the last half of the past decade, which, in most cases, reach double-digit figures (exceptions being the Czech Republic, Estonia and Latvia). For most countries, the initial transitional output collapse led to large increases in unemployment. This was especially so for Eastern European countries, whereas ex-Soviet republics experienced higher levels of labour hoarding and hence lower unemployment. The Czech Republic is the main exception to his pattern up until the stabilization plans of 1997 when

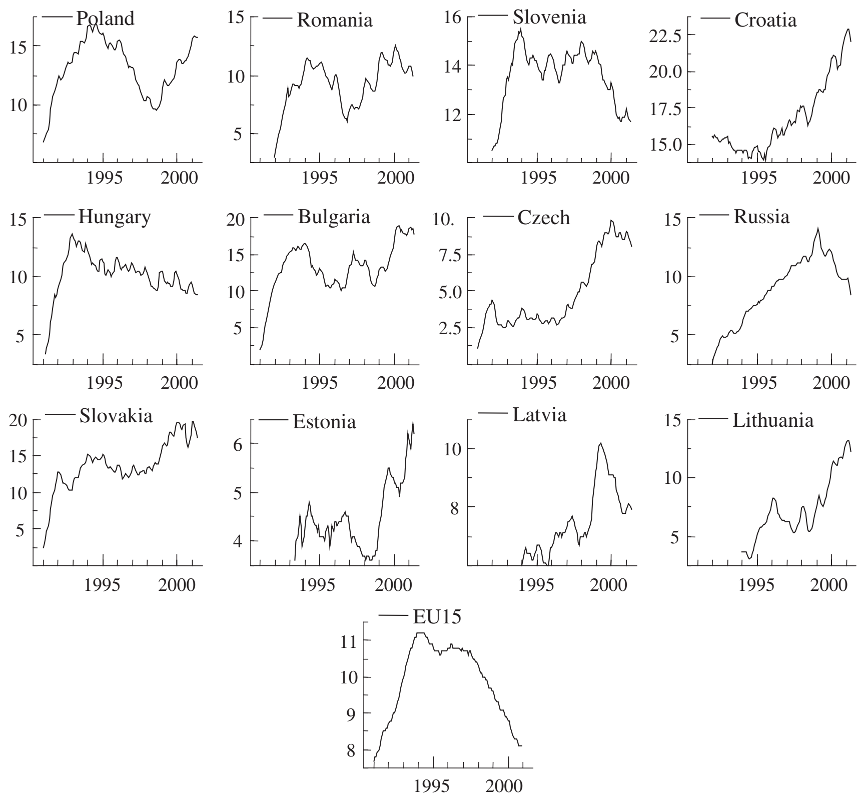


Figure 1. Unemployment rates.

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unemployment increased rapidly to around 10%. Countries such as Poland, Romania, Slovenia, Bulgaria and, only partially, Slovakia, experienced rapid increases followed by a recovery after 1994–95 and then a further deterioration. Estonia, Latvia and Lithuania all show initially low levels of unemployment that deteriorate rapidly by the end on the 1990s, especially for Lithuania. Russia shows a sharp increase in unemployment due to the 1998 crisis that starts to recover after 2000. Finally, Hungary is the only country for which we can find a steady recovery after the initial transformational shock.

Unemployment in these countries arose as a consequence of the rapid process of structural change and as the inevitable consequence of labour market reforms.[4] However, as Boeri and Terrell (2002) point out, more than the rate of employment destruction, it is the low rate of employment creation that has led to the existence of stagnant pools of long-term unemployment. This is especially the case in CEECs, whereas Russia and the CIS countries have shown consistently lower levels of unemployment at the expense of lower productivity levels. This happens even when the output collapse in former Soviet Union countries has been far larger than in most CEECs. This lower elasticity of employment with respect to output (essentially labour hoarding) is one of the main differences in employment performance between these two groups of countries.[5] The difference is related to the fact that wage adjustment has largely been thought a more prominent feature of labour-market dynamics in Russia, whereas employment has been the main adjustment variable in CEECs, pointing out to a higher degree of persistence of unemployment in the latter group.

These patterns of employment and unemployment dynamics are the consequence of a complex mix of events. Table 1 provides some summary statistics of labour participation, youth unemployment and educational levels of the labour force.[6] It is important to note that participation rates declined considerably. Indeed, as reported in Nesporova (2002) and Boeri and Terrell (2002), employment losses led partly to unemployment and partly to outflows from the labour force. That is, the market tensions were resolved by pushing seemingly less competitive groups out of the labour force. These outflows were considerably higher for the young (15–24) and for females. Female participation rates fell sharply for the Czech Republic, Slovakia, Latvia and Estonia. Although levels of unemployment are higher for females, this decline in participation makes the unemployment gender-gap comparable to those in Western economies. Youth unemployment, however, reaches the highest levels especially for Bulgaria, Slovakia and Poland. This is because demographic

[4] In the initial stages of transition, many countries pursued active labour policies such as wage subsidies, public-sector job creation, retraining, etc. as well as applying a range of income-support schemes. As unemployment rose in the immediate aftermath, however, these support schemes were restricted: the size, scope and maximum duration of unemployment benefits were typically reduced and wage subsidies cut. See Boeri (1997a, 1997b) for a review of labour market reforms in transition economies.

[5] Svejnar (1999) reports insignificant elasticities of employment to output for Russia and elasticities within the range of 0.2 and 0.8 for CEECs.

[6] See Nesporova (2002) and Rashid and Rutkowski (2001) for more detailed analyses of the microstructure of labour markets in transition.

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Table 1
*Labour market characteristics*

| | Participation rate by gender | | | Participation rate by age | | | Youth unempl. 2000 | Total unempl. 2000 | Level of education | | |
|---|---|---|---|---|---|---|---|---|---|---|---|
| | 1990 1999 | | | 1990 1999 | | | | | 2000 | | |
| | Men | Women | Total | 15–24 | 25–49 | 50–64 | | | Primary | Secondary | Tertiary |
| Poland | 80.1 | 65.1 | 72.5 | 44.3 | 87.3 | 60.6 | 35.7 | 16.6 | 33.1 | 58.3 | 8.6 |
| | 72.8 | 59.7 | 66.1 | 37.3 | 85.0 | 47.8 | | | | | |
| Romania | 76.7 | 60.5 | 68.5 | 59.8 | 87.7 | 42.9 | 17.8 | 7.7 | 43.2 | 49.9 | 6.9 |
| | 76.3 | 61.9 | 69.0 | 45.8 | 84.8 | 58.1 | | | | | |
| Slovenia | 76.7 | 64.8 | 70.7 | 50.4 | 93.1 | 42.0 | 16.4 | 7.1 | 33.9 | 53.9 | 12.1 |
| | 72.2 | 63.3 | 68.0 | 48.1 | 91.3 | 38.2 | | | | | |
| Croatia | 76.9 | 56.4 | 66.6 | 45.8 | 86.8 | 43.5 | n.a. | 21.1 | n.a. | n.a. | n.a. |
| | 59.8 | 51.4 | 55.6 | 40.3 | 70.5 | 23.3 | | | | | |
| Hungary | 74.5 | 57.3 | 65.4 | 51.5 | 86.0 | 36.0 | 12.3 | 6.6 | 38.5 | 50.3 | 11.2 |
| | 67.8 | 52.3 | 59.9 | 40.7 | 79.0 | 37.9 | | | | | |
| Bulgaria | 77.7 | 72.2 | 75.0 | 51.9 | 95.1 | 55.3 | 39.4 | 18.7 | 43.9 | 42.7 | 13.4 |
| | 75.9 | 64.9 | 70.2 | n.a. | n.a. | n.a. | | | | | |
| Czech Rep. | 82.2 | 74.1 | 78.1 | 57.7 | 96.0 | 55.7 | 17.0 | 8.8 | 23.8 | 67.0 | 9.1 |
| | 80.3 | 64.4 | 72.4 | 48.7 | 89.3 | 59.4 | | | | | |
| Slovak Rep. | 82.5 | 74.2 | 78.3 | 58.8 | 95.6 | 55.3 | 36.9 | 19.1 | 28.8 | 63.5 | 7.6 |
| | 76.1 | 64.2 | 69.3 | 45.6 | 89.5 | 45.6 | | | | | |
| Estonia | 83.3 | 75.9 | 79.4 | 53.0 | 95.6 | 68.5 | 23.7 | 13.5 | 26.2 | 51.3 | 22.5 |
| | 78.1 | 66.4 | 72.1 | 43.5 | 88.2 | 62.2 | | | | | |
| Latvia | 83.6 | 75.3 | 79.4 | 56.1 | 95.1 | 67.4 | 21.2 | 14.4 | 30.6 | 55.3 | 14.1 |
| | 75.3 | 62.6 | 68.7 | 41.6 | 87.0 | 53.1 | | | | | |
| Lithuania | 81.8 | 70.5 | 76.0 | 49.5 | 93.9 | 61.9 | 27.5 | 15.9 | 31.3 | 36.8 | 31.9 |
| | 77.4 | 68.3 | 72.7 | 39.8 | 92.6 | 59.8 | | | | | |
| Russia | 91.6 | 71.7 | 76.5 | 52.4 | 95.2 | 57.7 | n.a. | 10.5 | n.a. | n.a. | n.a. |
| | 74.2 | 63.9 | 68.9 | 41.9 | 87.7 | 48.8 | | | | | |

*Sources:* Nesporova (2002), EUROSTAT, Russian Economic trends, Statistical Office of Croatia.

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trends more than compensated for the lower participation rates. The lack of appropriate skills and seniority rules still existing in many privatized firms also contributed to the marked difference in unemployment rates of the young.

Another important characteristic of the labour force is its level of education. Although there is a wide variation amongst CEECs, the levels of educational attainment appear high for the level of development of these countries. Bulgaria and Romania show the highest share of primary education whereas the Czech Republic and Slovakia have a higher share of secondary education. Finally, the Baltic States show a high proportion of tertiary education population. However, these numbers hide the fact that the educational systems were very specialized and highly vocational (Boeri, 2001; Nesporova, 2002). This left many workers with obsolete skills unable to adapt to the rapidly changing demands of the private sector. This mismatch has led to the appearance of long-term unemployment accounting for more than 40% of total unemployment for most of these countries. Hence, the transitional unemployment of the initial phases of transformation has rapidly become structural in nature. This is of special relevance for our analysis.

In all, the unprecedented process of structural change that shook CEECs' labour markets has not been absorbed as expected by the creation of new jobs in the private sector and the improvement of matching induced by more market-oriented policies (EBRD, 2000). This has led to the high unemployment observed in the CEECs together with persistence and long duration of unemployment spells. However, as argued by Boeri and Terrell (2002) and Boeri (2001), it is difficult to associate this persistence pattern with the flexibility of labour markets. This is because the traditional factors used to explain maladjustment – e.g., union strength, minimum wages and employment protection legislation etc. – are either weak or effectively not implemented. For these authors, non-employment benefits[7] acting as wage floors may have discouraged job reallocation creating strong disincentive effects.

Theoretical models of multiple equilibria in transition labour markets have been developed by Aghion and Blanchard (1994), Garibaldi and Brixiova (1998) and Boeri (2001), amongst others. Aghion and Blanchard (1994) develop a model where, depending on agents' expectations, the transition economy could end up in a high unemployment equilibrium. In Boeri (2001), multiple equilibria can arise due to microeconomic lock-in effects, owing to the excessive skill specificity of workers together with the search disincentives generated by non-employment benefits in the formal and informal sectors.[8] This pattern would generate the appearance of long spells of unemployment and regime shifts. In many of these models, the timing of reforms determines the unemployment equilibrium (high or low). High persistence will arise even in effectively highly non-regulated labour markets such as those in CEECs. These models point to Hysteresis in unemployment. However, this mechanism substantially differs

[7] Mostly unemployment benefits, but also other forms of subsidies such as disability benefits.

[8] Garibaldi and Brixiova (1998) use a similar argument using labour market transition with a matching theoretical framework.

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from traditional models of persistence – such as Blanchard and Summers (1987) – based on insider-outsider effects or human capital loss.

If these theoretical models are correct, we should expect either high levels of persistence in unemployment dynamics or frequent unemployment equilibrium changes in the face of shocks such as those experienced by Eastern European countries. The first hypothesis has been traditionally tested on OECD countries by applying unit-roots tests to unemployment series, e.g. Song and Wu (1997) and Arestis and Mariscal (1999) and León-Ledesma (2002). The second hypothesis has been tested in Bianchi and Zoega (1998) and Jaeger and Parkinson (1994). Surprisingly, however, little effort has been made in studying unemployment dynamics in transition economies beyond mere descriptive analysis. We fill this gap by analyzing the persistence patterns of unemployment in Eastern Europe.[9]

At the outset, we should note that *Hysteresis*, though widely-used, carries many interpretations – see Røed (1997) for a survey – and indeed is often used interchangeably with *Persistence*. Hysteresis was originally used to imply that the 'natural rate' is dependent on the actual history of unemployment (Phelps, 1972); consequently, if a country suffers a prolonged period of historically-high unemployment, then equilibrium unemployment will itself rise, being, thus, both path dependent and non-unique. Persistence of unemployment, by contrast, implies a very slow adjustment of unemployment towards a unique natural rate – in statistical terms this translates into whether the series has a unit or a near unit root. If the root is high but below one we have 'partial hysteresis' and 'pure hysteresis' if the root is one (Layard *et al.*, 1991). In the latter case, equilibrium is not defined. Essentially, the unit-root definition of hysteresis is the one followed here, although we consider linear and nonlinear variants.

To illustrate, consider the following $AR(K)$ process for the unemployment rate ($y$):

$$y_t = \alpha_0 + \sum_{k=1}^{K} \alpha_k y_{t-k} + \varepsilon_t. \tag{1}$$

Here, the 'natural', mean or equilibrium rate to which unemployment reverts over time is $\bar{y} = (\alpha_0/(1 - \sum_k \alpha_k))$ – assuming $\sum_k \alpha_k < 1$ and no intercept shifts, i.e. $\alpha_0 = \alpha_0 \forall t$. However, if $\sum_k \alpha_k = 1$, unemployment follows a random walk and displays path-dependence (*pure* Hysteresis). Thus, shocks $\varepsilon_t$ have permanent effects.[10]

[9] By contrast, much of the debate over the chronic rise in European unemployment has focused less on the different shocks hitting the constituent economies and more on the interaction between these (often common) shocks and heterogeneous national institutional factors (e.g., Blanchard and Wolfers, 2000). Examining the interaction of shocks and institutions does not, however, originate with Blanchard and Wolfers; their analysis is an extension of Phelps (1994) ch. 17, which again is motivated by Layard *et al.* (1991). Shocks have a larger and more persistent effect in countries with 'poor' labour-market institutions. A suitable extension of our work, therefore, would be to consider the influence of such institutional factors in the evolution of labour markets in the transition economies. Needless to say, such an approach faces significant data constraints as well as the fact that the underlying institutions in transition countries have themselves been volatile.

[10] Note that, for the purposes of our analysis, these can be supply or demand shocks. It is beyond the scope of this paper to identify the relevant shocks.

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This is a particular concern for transition countries since (as in our previous discussion) it is not unreasonable to suppose that they have been hit by a relatively high number of shocks (increased openness to trade, price liberalization, privatizations and the removal of subsidies, changes in trading partners, etc.).

Testing for unit roots for the presence of pure linear Hysteresis provides an upper bound test of the hypothesis, given that this is an extreme case of path-dependence where any shock, large or small, matters. However, given that unemployment rates are necessarily bounded, unemployment should be stationary over longer time spans. Hence, Hysteresis as a unit root should not necessarily be understood as a 'true' description of the underlying data generating process but as a local approximation during a sample period. A less restrictive hypothesis considers Hysteresis as a process by which unemployment switches equilibria whenever 'sufficiently large' shocks affect its value; that is, if only large shocks enter the long-run memory of the unemployment series because they generate changes in the 'natural' or equilibrium level of unemployment.

Conventional stationarity tests can verify the presence of such 'unit roots'. However, testing for non-stationarity (in our application) raises a number of non-trivial technical issues. First, we necessarily have a short span of data. Second, tests may have low power against precisely those structural breaks that we might expect to characterize the data (e.g. the $\alpha_0$'s and $\alpha_k$'s may be time varying).[11] Third, if there are structural breaks, we must try to both date these and ensure that we distinguish them from normal business-cycle fluctuations. Finally, it is possible that unemployment takes – in contrast to equation (1) – some nonlinear form. This paper systematically tries to overcome these difficulties to robustly identify persistence patterns in transition countries' unemployment. On the first point (small sample), we use (in addition to conventional tests) panel unit-root tests that exploit both the time-series and cross-sectional dimensions of the data. As regards structural break tests (second and third points), we use single-equation and panel structural-break tests as well Markov-switching methods that endogenously search for and date structural breaks independent of normal cyclical fluctuations. Finally, our Markov-switching regressions control for any possible nonlinearity in the unemployment process and allow for the analysis of switching equilibrium unemployment as suggested both by theoretical models of labour markets in transition and, as already mentioned, by recent conceptualizations of unemployment Hysteresis.

## III Testing for Unit Roots

### *Time-series tests*

As mentioned, a traditional testing procedure is to apply unit root tests on the unemployment rate. The existence of a unit root would indeed imply that

[11] In our context, the most appealing form of break is an intercept break. This would be consistent with 'structural' explanations of the natural rate hypothesis. See Phelps (1994).

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unemployment does not revert to its natural rate after a shock. Table 2 presents four different but widely used unit root tests on the monthly, seasonally-adjusted unemployment series[12] of our set of 12 transition economies plus the EU-15 aggregate.[13] Details on data sources and sample periods can be found in the Appendix. The tests carried out are the Augmented Dickey-Fuller (ADF) test for the null of a unit root, the Kwiatkowski *et al.* (1992) LM test for the null of stationarity (KPSS) and the asymptotically most powerful DF-GLS tests of Elliot *et al.* (1996), ERS, and Elliott (1999) for the null of a unit root.[14] We report the tests with and without a time trend, and also provide the estimated auto-regressive root for the ADF test together with the derived half-life for the shocks. Given that our data is monthly, it is not surprising to observe high roots implying a slow speed of reversion to the mean. The results indicate that, for the majority of the tests, we cannot reject the null of a unit root for most countries in the sample. The main exception is Bulgaria, where only the ERS DF-GLS test for the model with an intercept cannot reject the null of a unit root. The other three countries where the presence of a unit root is rejected by several tests are Poland, Hungary and Lithuania and, to a lesser extent, Romania. On the other hand, Croatia, Estonia, Slovenia, Slovakia, Russia and the EU are shown to behave as unit root processes in most cases and, hence, have very large half-lives for the correction of shocks. For the Czech Republic, most tests including a time trend also reject the null of a unit root.

Confidence intervals for the largest auto-regressive root of the ADF tests were also constructed following Stock (1991). The 90% confidence intervals are reported in Table 3. Compatible with the previous results, only Romania and Bulgaria seem to lie within the unit interval. For the rest of the cases, the upper bound estimate of the largest root is higher than unity for at least one case. Note, however, that for countries such as Hungary, Czech Republic and Lithuania, the lower bound is sometimes close to 0.6, implying a very fast adjustment to shocks with around 1 month half-life.[15] Another aspect of relevance is that the confidence intervals are, with a few exceptions, reasonably tight given our short sample and number of observations.

The tests presented in Table 2, however, suffer from two important problems that could substantially reduce their reliability. First, as pointed out by Perron (1989), in the presence of a structural change, we could erroneously be favouring

[12] We use high-frequency (i.e., monthly) data to capture the aspects of the rapidly changing CEEC's labour markets. In addition, monthly as opposed to quarterly data, provides degrees of freedom that improve (or at least do not deteriorate) the precision of our estimates (particularly in our Markov Switching regressions).

[13] We also performed our tests with an EU-12 aggregate with little change in our results (details available).

[14] The main difference between the ERS and the Elliott (1999) tests is that the former assumes zero initial conditions for the process under both the null and alternative, while the latter draws the initial observation from its unconditional distribution under the alternative. It is not our purpose to discriminate between these various stationarity tests in terms of power or size (e.g. Caner and Kilian, 2001).

[15] The 0.0% confidence interval does not, in general, coincide with the point estimate in Table 2 because, as argued by Stock (1991), the local-to-unity distribution of the point estimate of the auto-regressive root is skewed and depends on nuisance parameters.

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Table 2
*Unit root tests*

| | | ADF | | | | KPSS | | ERS | | Elliott (1999) | |
|---|---|---|---|---|---|---|---|---|---|---|---|
| | | Intercept model | | Trend and intercept model | | Intercept model | Trend model | Intercept model | Trend model | Intercept model | Trend model |
| Country series | No. of Lags | *T*-ratio | Estimated root {half life} | *T*-ratio | Estimated root {half life} | $\eta_\mu$ | $\eta_\tau$ | DF-GLS | DF-GLS | DF-GLS$_u$ | DF-GLS$_u$ |
| Poland | 12 | – 2.848 | 0.973 {25.32} | – 2.363 | 0.977 {29.79} | 0.1518* | 0.1580 | – 2.432 | – 3.206* | – 2.801* | – 3.214* |
| Romania | 10 | – 3.386* | 0.945 {15.05} | – 3.648* | 0.947 {13.80} | 0.3210* | 0.1116 | – 2.120 | – 1.895 | – 1.927 | – 2.070 |
| Slovenia | 11 | – 2.324 | 0.944 {12.03} | – 2.752 | 0.928 {9.28} | 0.1904* | 0.1820 | – 1.291 | – 1.367 | – 1.649 | – 1.367 |
| Croatia | 12 | 0.548 | 1.006 {n.a.} | – 2.348 | 0.941 {11.40} | 0.8212 | 0.2373 | – 0.694 | – 1.345 | – 0.611 | – 1.821 |
| Hungary | 12 | – 2.118 | 0.950 {13.51} | – 4.527* | 0.841 {4.00} | 0.2112* | 0.1398* | – 2.260 | – 2.921* | – 3.059* | – 2.901 |
| Bulgaria | 12 | – 3.523* | 0.947 {12.73} | – 4.379* | 0.928 {9.28} | 0.3790* | 0.1080* | – 2.812* | – 3.195* | – 3.464* | – 3.846* |
| Czech Rep. | 12 | – 0.897 | 0.995 {138.28} | – 4.697* | 0.949 {13.24} | 0.8479 | 0.2218 | – 3.118* | – 4.172* | – 0.912 | – 4.109* |
| Slovak Rep. | 12 | – 0.782 | 0.988 {57.41} | – 2.054 | 0.947 {12.72} | 0.7864 | 0.1201* | – 1.893 | – 2.226 | – 0.501 | – 2.251 |
| Estonia | 7 | – 0.932 | 0.971 {23.55} | – 1.616 | 0.966 {20.04} | 0.5896 | 0.2313 | – 1.892 | – 1.892 | – 1.191 | – 1.885 |
| Latvia | 2 | – 1.638 | 0.986 {49.16} | – 2.283 | 0.950 {13.55} | 2.0493 | 0.1866 | – 2.277 | – 2.385 | – 1.620 | – 2.364 |
| Lithuania | 12 | – 2.001 | 0.974 {26.31} | – 4.673* | 0.897 {6.38} | 0.6341 | 0.1324* | – 2.533 | – 4.188* | – 2.033 | – 4.112* |
| Russia | 12 | – 1.719 | 0.989 {62.67} | – 1.431 | 0.982 {39.16} | 0.8217 | 0.2013 | – 1.187 | – 1.312 | – 1.533 | – 1.461 |
| EU-15 | 6 | – 1.687 | 0.989 {62.67} | – 2.0411 | 0.986 {49.16} | 0.4285* | 0.4215 | – 1.576 | – 1.716 | – 1.898 | – 1.891 |

*Notes:*
*Indicates rejection of the null of a unit root at the 5% level for the ADF, DF-GLS and DF-GLS$_u$ tests respectively and not rejection of the null of stationarity for the KPSS test at the 95% level.
The half-life was calculated as – Ln(2)/Ln($\varsigma$), where $\varsigma$ is the auto-regressive root of unemployment in the ADF test, and is expressed in months.
The number of lags was chosen using a General-to-Specific criterion.

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Table 3
*90% Confidence intervals for the auto-regressive root in ADF test*

| | Intercept model | | Trend model | |
|---|---|---|---|---|
| Country series | 90% interval | 0.0% interval {half life} | 90% interval | 0.0% interval {half life} |
| Poland | (0.812, 1.001) | 0.895 {6.25} | (0.879, 1.028) | 0.969 {22.01} |
| Romania | (0.737, 0.935) | 0.833 {3.79} | (0.726, 0.956) | 0.832 {3.77} |
| Slovenia | (0.867, 1,015) | 0.940 {11.20} | (0.838, 1.023) | 0.928 {9.28} |
| Croatia | (1.004, 1.034) | 1.016 {n.a.} | (0.880, 1.028) | 0.971 {23.55} |
| Hungary | (0.886, 1.019) | 0.956 {15.40} | (0.592, 0.841) | 0.710 {2.02} |
| Bulgaria | (0.729, 0.929) | 0.827 {3.65} | (0.624, 0.869) | 0.741 {2.31} |
| Czech Rep. | (0.972, 1.030) | 1.010 {n.a.} | (0.579, 0.829) | 0.698 {1.93} |
| Slovak Rep. | (0.978, 1.031) | 1.011 {n.a.} | (0.908, 1.031) | 1.013 {n.a.} |
| Estonia | (0.970, 1.030) | 1.010 {n.a.} | (0.946, 1.034) | 1.017 {n.a.} |
| Latvia | (0.926, 1.025) | 0.996 {172.94} | (0.887, 1.029) | 0.979 {32.66} |
| Lithuania | (0.896, 1.020) | 0.965 {19.45} | (0.583, 0.832) | 0.701 {1.95} |
| Russia | (0.942, 1.027) | 1.005 {n.a.} | (1.017, 1.041) | 1.024 {n.a.} |
| EU-15 | (0.919, 1.025) | 0.992 {86.30} | (0.905, 1.033) | 1.014 {n.a.} |

*Note:*
Confidence intervals calculated using Stock's (1991) method.

the existence of a unit root when the process is in fact stationary with a change of mean or trend.[16] The second problem is the low power of these tests especially when the sample is small. Although we are dealing with series of around 120 observations, our sample period of about 10 years might reduce the power of our tests and, hence, bias the results towards the acceptance of the null of a unit root. We attempt to deal with the latter shortcoming later when making use of panel unit root tests.[17]

In order to control for the presence of structural breaks on the ADF regressions, we carried out the Perron (1997) unit-root test with endogenous search for structural change, based on the following ADF regression for time series $y_t$:

$$y_t = \alpha_0 + \theta DU_t + \beta t + \delta D(T_b) + \alpha_1 y_{t-1} + \sum_{k=1}^{K} \Delta y_{t-k} + \varepsilon_t \qquad (2)$$

[16] To illustrate possible instability of the ADF regressions and the existence of structural breaks, we also obtained recursive Chow stability tests of the auto-regressive form of the ADF test, $AR(K)$, with $K$ being the maximum lag chosen for the unit root tests. The Chow test exceeds its 5% critical value for several countries (details available in our working paper version, León-Ledesma and McAdam, 2003).

[17] We also checked for the possibility of an asymmetric adjustment of unemployment in expansion and slowdown periods by fitting a momentum threshold auto-regressive model (M-TAR) to our data. The results did not show significant asymmetries in unemployment dynamics except for the possible case of the Czech Republic, which showed a higher persistence in periods of unemployment reduction. This, however, did not change our previous conclusions about unit roots in this country (details available on request).

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where $DU_t = 1$ $(t > T_b)$ and $D(T_b)_t = 1$ $(t = T_b+1)$ with $T_b$ being the time at which the change in the trend function occurs, and $K$ is the lag augmentation for correction of residual auto-correlation. This is Perron's (1989) 'innovational outlier model' that implies a change in the mean but not the slope of the ADF regression. This is the most likely case to occur in unemployment series because of changes in the 'natural' rate. The test for a unit root is performed using the $t$-statistic for the null hypothesis that $\alpha_1 = 1$. The optimal search for the break date is carried out using two methods. The first finds $T_b$ as the value that minimizes the $t$-statistic for testing $\alpha_1 = 1$. In the second, $T_b$ is chosen to maximize the absolute value of the $t$-statistic associated with the change in the intercept $\theta$.[18] As is standard in structural break tests, we have limited the search of the break date for both methods excluding the first and last 10% sample observations.

Table 4 reports the results of the Perron (1997) test. We report the break date ($T_b$), the $t$-statistic for $\alpha_1 = 1$ and the estimated auto-regressive root for both break search methods. The $t$-statistics are compared with the critical values for $T = 100$ provided by Perron (1997). For 10 out of 13 economies tested, both methods gave the same break date (or very similar in the case of Slovakia).[19] The results for the unit-root test indicate that we can now reject the null of non-stationarity for 6 countries by at least one of the methods. The speed of adjustment is now substantially faster in all cases as reflected in lower values of the estimated root. Of our sample, only Poland gets close to the EU aggregate in terms of the calculated half lives. For some countries like Hungary or Russia, the half-life becomes close to 3 months. Thus, once we have controlled for structural breaks, Hysteresis in our set of transition countries appears less plausible.

### *Panel tests*

As mentioned earlier, because of our relatively short sample, traditional unit root tests may suffer a lack of power. To solve this, several authors have proposed the use of panel unit root tests that exploit both the time-series and cross-sectional dimensions of the data. Here we apply three such tests; two – Im *et al.* (2003) and Chang (2002) – rely on panel versions of ADF regressions whilst the third, Taylor and Sarno (1998), is based on Johansen's (1992) Likelihood Ratio test for cointegration in a VAR.[20]

[18] We chose this method instead of minimizing the $t$-statistic on $\theta$ to avoid imposing *a priori* assumptions on the sign of the change.

[19] An interesting feature is that the two break methods tend to give more breaks in the second half of the 1990s. This is compatible with labour-market research in CEECs that emphasizes the deterioration of unemployment around 1997–1999 and, as will be discussed in Section IV, with our results from Markov-switching regressions.

[20] Note that the Panel tests used are all designed for heterogeneous panels in which each cross-section is estimated independently and not pooled; the IPS, Chang and Taylor and Sarno tests do not impose the same speed of mean reversion of unemployment rates in these countries. Hence, convergence in unemployment is not necessary to make use of these tests. Thus, we have not used tests with homogeneous panels such as Levin *et al.* (2002), although results based on this latter test are available on request. This seems an appropriate choice as visual inspection of unemployment figures reveals a high degree of heterogeneity in the sample.

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Table 4
*Perron (1997) tests on unemployment series*

| | | Break search model I | | | Break search model II | | |
|---|---|---|---|---|---|---|---|
| Country series | Period | Break date | *T*-ratio | Estimated root {half life} | Break date | *T*-ratio | Estimated root {half life} |
| Poland | 91:01–01:05 | 1996:03 | – 4.372 | 0.946 {13.49} | 1996:03 | – 4.372 | 0.946 {13.49} |
| Romania | 91:12–01:04 | 1993:11 | – 5.032** | 0.915 {8.80} | 1996:01 | – 4.537 | 0.880 {6.42} |
| Slovenia | 91:12–01:04 | 1999:06 | – 4.578 | 0.835 {4.84} | 1999:06 | – 4.578 | 0.835 {4.84} |
| Croatia | 91:01–01:05 | 1999:01 | – 4.017 | 0.880 {6.42} | 1999:01 | – 4.017 | 0.880 {6.42} |
| Hungary | 91:05–01:08 | 1992:11 | – 5.947* | 0.680 {2.80} | 1992:11 | – 5.947* | 0.680 {2.80} |
| Bulgaria | 91:01–01:05 | 1999:02 | – 5.28* | 0.907 {8.10} | 1999:02 | – 5.28* | 0.907 {8.10} |
| Czech Republic | 91:01–01:05 | 1998:04 | – 6.781* | 0.891 {7.01} | 1992:07 | – 3.789 | 0.961 {18.42} |
| Slovak Republic | 91:01–01:05 | 1992:11 | – 3.580 | 0.908 {8.18} | 1992:12 | – 3.477 | 0.909 {8.26} |
| Estonia | 93:05–01:05 | 2000:05 | – 3.431 | 0.912 {8.52} | 1998:10 | – 2.343 | 0.911 {8.43} |
| Latvia | 94:01–01:05 | 1998:04 | – 4.195 | 0.850 {5.26} | 1998:04 | – 4.195 | 0.850 {5.26} |
| Lithuania | 94:01–01:05 | 1997:01 | – 8.153* | 0.791 {3.96} | 1997:01 | – 8.153* | 0.791 {3.96} |
| Russia | 92:01–01:04 | 1998:08 | – 6.379* | 0.732 {3.22} | 1998:08 | – 6.379* | 0.732 {3.22} |
| EU-15 | 91:01–00:12 | 1992:05 | – 3.800 | 0.967 {21.66} | 1992:05 | – 3.800 | 0.967 {21.66} |

*Notes:*
*and **indicate rejection of the null at the 5% and 10% significance level respectively.
Critical values from Perron (1997) Table 1.
Half-life calculated as in Table 2.

The Im *et al.* (2003) (IPS) test is based on the ADF regression:[21]

$$\Delta y_{i,t} = \alpha_{0i} + \varsigma_i y_{i,t-1} + \sum_{k=1}^{K_i} \gamma_{ik} \Delta y_{i,t-k} + \varepsilon_{i,t} \tag{3}$$

where $i = 1, 2, \ldots, N$ and $t = 1, 2, \ldots, T$. (The) IPS test(s) the null of an I(1) process ($\varsigma_i = 0\ \forall i$) against the alternatives $H_A$: $\varsigma_i < 0$, $i = 1, 2, \ldots, N_1$, $\varsigma_i = 0$, $i = N_1+1, N_1+2, \ldots, N$. Note that the IPS test does not assume that all cross-sectional units converge towards the equilibrium value at the same speed under the alternative, i.e. $\varsigma_1 = \varsigma_2 = \ldots = \varsigma_N < 0$, and thus is a less restrictive test than previous panel tests such as Levin *et al.* (2002), LLC. The IPS test is based on the

[21] For simplicity, we ignore deterministic trends in the explanation of the tests.

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standardized $t$-bar statistic:

$$\Gamma_t = \frac{\sqrt{N}[\bar{t}_{NT} - \mu]}{\sqrt{v}} \sim N(0, 1) \tag{4}$$

where $\bar{t}_{NT}$ is the average of the $N$ cross-section ADF($K_i$) $t$-statistics. $\mu$ and $v$ are, respectively, the mean and variance of the average ADF($K_i$) statistic under the null, tabulated by Im *et al.* (2003) for different $T$s and lag orders of the ADF. Im *et al.* (2003) also show that under the null of a unit root, $\Gamma_t$ converges to an $N(0, 1)$ as $N/T \to q$ (where $q$ is any finite positive constant).

One of the problems of the IPS test is that it assumes that the different cross-sections are distributed independently. One way to avoid this, as suggested by Im *et al.* (2003), is to subtract cross-sectional averages from the individual series. This, however, does not allow for more general forms of dependency. The test proposed by Chang (2002) tries to overcome this problem by using a non-linear IV estimation of the individual ADF regressions using as instruments non-linear transformations of the lagged levels. The standardized sum of individual IV $t$-ratios has a limit normal distribution. Here we used the following Instrument Generating Function as in Chang (2002):

$$F(y_{i,t-1}) = y_{i,t-1} e^{-c_i |y_{i,t-1}|}. \tag{5}$$

Where $c_i$ is proportional to the sample standard error ($\sigma$) of the first difference of $y_{it}$:

$$c_i = JT_i^{-1/2} \sigma_{\Delta y_{it}} \tag{6}$$

where $J$ is a constant fixed at 5 as recommended in Chang (2002) for time dimensions larger than 25 observations.

The Taylor and Sarno (1998) test (TS) takes a different route based on Johansen's (1992) Maximum Likelihood method to determine the number of common trends in a system of unit-root variables. We can represent an $N$-dimensional vector auto-regressive (VAR) process of $K^{\text{th}}$ order as:

$$\Delta Y_t = \alpha + \Theta_1 \Delta Y_{t-1} + \cdots + \Theta_{K-1} \Delta Y_{t-K+1} + \Pi Y_{t-K} + \varepsilon_t, \tag{7}$$

where $\alpha$ is a ($N \times 1$) vector of constants, $Y_t$ is a ($N \times 1$) vector of time series, $\Theta_{\mathrm{i}}$ are ($K \times K$) matrices of parameters and $\Pi$ is a ($N \times N$) matrix of parameters whose rank contains information about long-run relationships between the variables in the VAR. If $\Pi$ has full rank ($rank(\Pi) = N$) then all variables in the system are stationary. Hence, the TS test has as a null $H_o$: $rank(\Pi) < N$ and as alternative $H_A$: $rank(\Pi) = N$, which can be implemented using Johansen's (1992) Likelihood Ratio test. That is, it tests if one or more of the system variables is non-stationary against the alternative that all the variables are stationary. This is a more restrictive test than LLC and IPS because it will reject the null *if at least one* of the series in the panel has a unit root.

The results from these three tests are presented in Table 5. We have carried out the test for three different groups. The first one contains all the transition economies. The second excludes Bulgaria, since this was the only economy in

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Table 5
*Panel unit-root tests results on CEECs*

| | | Im, Pesaran and Shin (1997) | Chang (2002) | Taylor and Sarno (1998) |
|---|---|---|---|---|
| Unadjusted 12 countries | Intercept | – 0.962 | – 1.636** | 11.490* |
| | Trend | – 3.809* | – 2.795* | 14.750* |
| Adjusted 12 countries | Intercept | – 2.202* | – | – |
| | Trend | – 3.381* | – | – |
| Unadjusted 11 countries | Intercept | – 0.461 | – 1.404** | 11.955* |
| | Trend | – 3.327* | – 2.663* | 13.122* |
| Adjusted 11 countries | Intercept | – 1.698* | – | – |
| | Trend | – 2.738* | – | – |
| Unadjusted 9 countries | Intercept | – 4.245* | – 1.711* | 7.891* |
| | Trend | – 3.931* | – 2.754* | 10.147* |
| Adjusted 9 countries | Intercept | – 3.510* | – | – |
| | Trend | – 2.855* | – | – |

*Notes:*
Estimation periods for 12 and 11 countries are 1994M1–2001M4.
Estimation period for 9 countries is 1992M1–2001M4.
12 countries include all the database.
11 countries excludes Bulgaria, which was shown not to have a unit root in ADF tests.
9 countries exclude Estonia, Latvia and Lithuania due to their shorter time series.
For the IPS and Chang tests we used the lags chosen from the ADF tests. We used 4 lags for TS test because of lack of degrees of freedom to estimate a larger lag structure in a sensible way. However, the TS test results are not sensible to the inclusion of up to 6 lags. The critical values for the TS test ($\chi^2(1)$) have been adjusted by a factor $T/(T-K{\cdot}N)$ as recommended by Taylor and Sarno (1998), where $K$ is the lag of the VAR and $N$ is the number of countries. Results for the TS test with adjusted data are not possible to obtain, because the adjustment method would obviously lead to multicollinearity in the VAR.
Results for the Chang test with adjusted data are not reported given that the test controls for cross-sectional dependence.
*and **indicate rejection of the null of a unit root at the 5% and 10% level respectively.

which nearly all time series tests rejected non-stationarity. Given that the null of the IPS and Chang (2002) tests is that *all* cross-sections have a unit root, they would clearly be affected by the inclusion of a stationary series. The third group contains all economies except Estonia, Latvia and Lithuania for which data only starts in 1994M1, and shortens the time-series component of the panel. As the IPS test loses power if there is substantial cross-sectional correlation in the panel, we also applied the tests to each series adjusted by subtracting the cross-sectional average. Overall, the results show that the unemployment series are stationary. Only the null of the test on unadjusted data and an intercept for the 11 countries group seems to indicate the presence of a unit root. The IPS test rejects the null in all cases but two and the Chang (2002) and TS test, the most restrictive, reject the null in all cases.[22]

Finally, given the evidence on the likely importance of structural breaks, we combine panel unit root tests with endogenous break search tests by using the Murray and Papell (2000) (MP) test. This test can be considered a combination

[22] We compared the TS test to a $\chi^2(1)$ adjusted by a factor $T/(T\text{–}K{\cdot}N)$ as recommended by Taylor and Sarno (1998) to account for finite sample bias.

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Table 6
*Murray-Papell break panel unit root test*

| | Period | Break date | *T*-ratio | Estimated root {half-life} |
|---|---|---|---|---|
| 12 countries | 1994:01–2001:04 | 1998:09 | – 12.115* | 0.899 {6.51} |
| 11 countries | 1994:01–2001:04 | 1998:05 | – 10.065* | 0.920 {8.31} |
| 9 countries | 1991:01–2001:04 | 1996:03 | – 9.978* | 0.939 {11.01} |

*Notes*:
11 countries includes all countries in the sample except Bulgaria.
9 countries excludes Estonia, Latvia and Lithuania.
The critical values are given by Murray and Papell (2000). For $N = 10$ and $T = 100$, the 1% critical value is $-8.658$ and for $T = 50$ and $N = 10$ it is $-9.056$.
*indicates rejection of the null at the 5% level.

of the Perron (1997) test and the Levin *et al.* (2002) panel unit root test. It assumes that the auto-regressive coefficient of all cross-sections is the same and that the date of break is common between cross-sections. It allows for heterogeneity in the intercept and the lag augmentation of the ADF equation and accounts for cross-sectional correlation by estimating the panel by SUR methods. The break date is found as the one that minimizes the *t*-statistic for testing $\alpha_1 = 1$ as in Perron's (1997) Method I. The results of this test are reported in Table 6. We chose the lag augmentation of each unit to be the same found for the ADF tests and, again, applied the test for the 3 groups considered in previous panel tests. The results, again, strongly reject the null of non-stationarity at the 1% level, and the auto-regressive roots are found to be of the order of 0.9. The dating of the break in the second half of the 1990s is not surprising, given the rapid deterioration of unemployment in many countries during this period and the results obtained using the Perron (1997) test for individual countries.

The overall picture suggests that unemployment dynamics in transition economies during the last decade have not been characterized by a linearly hysteretic behaviour. Once we control for the impact of structural change, the low power of time series tests, or both, we can reject a random walk in unemployment. Although there is still a high level of persistence in countries such as Croatia, Slovenia, Estonia or Latvia, on average, it is lower than that for the EU aggregate. The lock-in effects that theory models describe at the micro level do not appear to have derived from a random walk behaviour.

## IV Markov-Switching Analysis

So far, our analysis has been confined to testing for a strong version of Hysteresis that assumes that every shock has permanent effects on the level of unemployment. However, as mentioned earlier, Hysteresis has also been associated with the existence of multiple equilibria in unemployment dynamics, e.g. Amable *et al.* (1995). Importantly, our previous analysis of unit roots makes a number of assumptions, which we might now like to relax or reconsider. First, the unit-root structural-break tests which, being essentially supremum tests,

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might be considered as biased towards finding a break even if one does not exist. Secondly, this is particularly problematic if the data (as we might suppose) is characterized by both business-cycle fluctuations and possibly, structural breaks. Third, the break implicit in the analysis of unit roots of the previous section assumes that either unemployment reverts to a constant level or to an average characterized by sudden changes. Unemployment, however, is more likely to adapt more smoothly to an infrequently changing average or 'natural' level of unemployment. That is, that unemployment is subject to changes in regimes.

These regime changes in the case of CEECs may be due to the multiple equilibria features arising from theoretical models described in Section II. Given the specificities of labour (and goods) markets of these countries, lock-in effects due to microeconomic factors may be important as these economies suffer large shocks stemming from the rapid pace of reforms. Such nonlinear behaviour has also been incorporated in empirical models of the Phillips curve – e.g., Gruen *et al.* (1999) – to reflect 'speed-limit' effects on the NAIRU. For these reasons, and to add an extra layer of robustness to our results, we analyze unemployment persistence using Markov-switching regressions. This will allow us not only to test for Hysteresis with a changing average level of unemployment, but also to analyze the frequency of regime changes and the behaviour of unemployment in each regime. Another advantage is that it accounts for nonlinearities in the trend unemployment function accruing not only from structural breaks but also from normal business-cycle fluctuations.

Our Markov-switching (MS) model for $m$ regimes (or states), $m \in [2, \infty)$, can be represented by equation (8) where $y_t$ (the unemployment rate) is regressed on an intercept ($\alpha_0$), auto-regression of length $K$ and a residual ($\varepsilon_t$) with variance $\sigma^2(s_t)$, all of which might be state dependent (denoted by $s_t$):[23]

$$y_t = \alpha_0(s_t) + \sum_{k=1}^{K} \alpha_k(s_t) y_{t-k} + \varepsilon_t,\ s_t = 1, \ldots, m. \tag{8}$$

The notable characteristic of such models is the assumption that the unobservable realization of the state, $s_t$, is governed by a discrete-time, discrete-state Markov-stochastic process. This is defined by the transition probabilities:

$$\Pr(s_{t+1} = j | s_t = i) = \rho_{ij},\ \sum_{j=1}^{m} \rho_{ij} = 1,\ \forall i. \tag{9}$$

[23] An alternative to the intercept-switching model (8) is the switching-in-mean ($\mu$) model:

$$y_t = \mu(s_t) + \sum_{j=1}^{J} \alpha_j [y_{t-j} - \mu(s_{t-j})] + u_t,$$

which is popular in, for instance, business-cycle analysis (where $y$ is the real growth rate) and financial modelling. In that model, after a change in the state there is an immediate one-time jump in the mean process. Since we are dealing with a labour market rather than, say, a spot financial market, we consider it more plausible that the mean should slowly and gradually adjust to a new level (from one transition to another) rather than as an immediate mean jump. The data also strongly suggested intercept over mean dependency (details available on request).

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Thus, $s_t$ follows a Markov process with the transition probabilities matrix, $P$:

$$P = \begin{bmatrix} \rho_{11} & \rho_{12} & \cdots & \rho_{1m} \\ \rho_{21} & \rho_{22} & \cdots & \rho_{2m} \\ \vdots & \vdots & \ddots & \vdots \\ \rho_{m1} & \rho_{m2} & \cdots & \rho_{mm} \end{bmatrix}. \tag{10}$$

Defining the number of states ($m$) is among the most difficult aspects of MS models (e.g. Garcia and Perron, 1996). We used two identification methods: (1) choosing $m$ based on Likelihood criteria, (2) using kernel density estimation methods, we use the number of modes in the density function as an indicator for the number of states.[24]

We estimate using the EM algorithm (Hamilton, 1990) and assign an individual observation $y_t$ to the state $m$ with the highest 'smoothed' probability:

$$m^* = \arg\max_m \Pr(s_t = m | y_T, y_{T-1}, \ldots, y_1).$$

In addition, we use a parametric bootstrap procedure whereby errors are repeatedly drawn from a $N(0, \hat{\sigma}^2)$ distribution and the model re-estimated to, for instance, derive standard errors for the composite parameter $\sum_k \alpha_k$. Furthermore, in terms of inference, we follow Hall *et al.* (1999) and Nelson *et al.* (2001) who conclude that DF/ADF tests remain robust in detecting stationarity in Markov-Switching regressions. Finally, Likelihood-Ratio tests confirm state-dependent variances.

Given our sample coverage, we find essentially only two states in the data. Exceptions are Poland, Romania and Croatia, for whom we model one.[25] Table 7 presents country estimates of the summed auto-regressive parameter $\sum_k \alpha_k$, state error variances $\sigma_i^2$, state means, $\bar{y}_i$, transition probabilities $\rho_{ii}$ and proportion measures $\xi_i$. First of all, we see that – excluding Latvia and second-state Czech Republic – having controlled for different states (i.e. business-cycle fluctuations and/or structural breaks) all countries have stationary processes for their unemployment rates. As before, the country with the highest level of persistence – and thus the slowest adjustment to a shock – is the EU-15. In many cases, we see that there has been a rather unbalanced state dependence. For example, most countries (excepting Bulgaria and the Czech Republic) spend around two thirds of their time in one state. States are also highly duration persistent: once in state $i$, the probability of remaining there is around 0.8 and upwards. (The exception appears to be Hungary where there has been considerably more switching between states.) Notably, in those cases where

[24] These two state identification methods are discussed more fully in León-Ledesma and McAdam (2003). The first method relies on Psaradakis and Spagnolo (2003). The second on the multi-modality bootstrap literature, Silverman (1986) and Efron and Tibshirani (1993).

[25] Thus, for these single-state cases, the stationarity tests already reported remain the measure of Hysteresis.

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Table 7
*Markov-switching results*

| Country series | $A_1(L)^{(1)}$ | $A_2(L)$ | $\sigma_1^2 = \sigma_2^2$ | $\sigma_{u,1}^2$ | $\sigma_{u,2}^2$ | $\bar{y}_1{}^{(2)}$ | $\bar{y}_2$ | $\rho_{11}{}^{(3,4)}$ | $\rho_{22}$ | $\xi_1{}^{(4)}$ | $\xi_2$ |
|---|---|---|---|---|---|---|---|---|---|---|---|
| Slovenia | 0.8226 | | 9.6762 | 0.004 | 0.018 | 12.215 | 14.196 | 0.829 | 0.942 | 0.255 | 0.745 |
| | (0.0367) | | [0.002] | (0.0079) | (0.007) | | | | | | |
| Hungary | 0.939 | | 42.5960 | 0.009 | 0.135 | 9.072 | 11.742 | 0.767 | 0.560 | 0.653 | 0.347 |
| | (0.003) | | [0.000] | (0.005) | (0.0066) | | | | | | |
| Bulgaria | 0.9015 | | 1.3212 | 0.072 | | 11.580 | 15.783 | 0.942 | 0.947 | 0.478 | 0.522 |
| | (0.014) | | [0.250] | (0.002) | | | | | | | |
| Czech Republic | 0.690 | 0.979 | 2.2744 | 0.009 | | 3.028 | 6.786 | 0.980 | 0.982 | 0.467 | 0.533 |
| | (0.152) | (0.141) | [0.132] | (0.001) | | | | | | | |
| Slovak Republic | 0.9293 | | 16.672 | 0.060 | 0.263 | 13.314 | 17.902 | 0.987 | 1.000 | 0.677 | 0.323 |
| | (0.0264) | | [0.000] | (0.028) | (0.028) | | | | | | |
| Estonia | 0.937 | | 7.296 | 0.004 | 0.0356 | 3.803 | 5.931 | 0.905 | 0.882 | 0.556 | 0.444 |
| | (0.015) | | [0.0069] | (0.0017) | (0.002) | | | | | | |
| Latvia | 0.935 | 0.824 | 12.240 | 0.0126 | 0.0027 | 7.119 | 9.304 | 0.982 | 0.935 | 0.785 | 0.215 |
| | (0.099) | (0.323) | [0.001] | (0.0034) | (0.00036) | | | | | | |
| Lithuania | 0.849 | 0.906 | 4.388 | 0.0228 | 0.011 | 6.226 | 8.938 | 0.978 | 1.000 | 0.595 | 0.405 |
| | (0.008) | (0.008) | [0.036] | (0.0001) | (0.0001) | | | | | | |
| Russia | 0.979 | | 27.306 | 0.0003 | 0.0315 | 9.917 | 9.509 | 0.808 | 0.925 | 0.280 | 0.720 |
| | (0.004) | | [0.000] | (0.0005) | (0.0005) | | | | | | |
| EU-15 | 0.987 | | 0.001 | 0.0029 | | 9.129 | 10.112 | 0.952 | 0.855 | 0.751 | 0.249 |
| | (0.005) | | [0.981] | (0.0002) | | | | | | | |

*Notes:*
(1)$A_i(L) = \sum_k \alpha_k | m = i$.
(2)Mean unemployment rate in $i^{\text{th}}$ state: $\bar{y}_i$.
(3)$\rho_{ii} = \Pr(s_{t+1} = i | s_t = i)$.
(4)Proportion of time in $i^{\text{th}}$ state: $\xi_i = \frac{n_{t,i}}{T}$, $\sum_{i=1}^{m} \xi_i = 1$ where $T$ = sample size, $n_{t,i}$ = number of observations in $i^{\text{th}}$ state and, $\sum_t \sum_{i=1}^{m} n_{t,i} = T$ Standard errors in ()'s, $p$-values in []'s. In each case, there are as many intercepts as states.

there exists state-dependent variance, high unemployment generally accords with high variance.

The country with the highest effective level of persistence statistically is the Czech Republic. This is because we cannot reject the null of a unit root in state 2.[26] However, state one, where the Czech economy spends nearly 47% of the observations, presents a very low auto-regressive root. These two states identify the rapid process of labour-market deterioration suffered by the Czech economy during the late 1990s. Another important result is that, for the majority of cases, and in line with previous unit-root tests, we reject the null of a random-walk behaviour. Furthermore, unemployment mean rates across states appear relatively well separated (e.g., the Czech Republic has an average unemployment rate of 3.0% in state 1 and 6.8% in state 2). With the exception of the EU-15 and Russia – where the spread is marginal – our results lend strong support to the notion of multiple equilibria.

States captured by MS methods can be both business-cycle fluctuations (recessions and expansions) as well as structural breaks. A concept related to

[26] The MS model also suggests a unit root in the case of Latvia, although this derives more from the imprecision of the standard errors than a high point value.

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the latter is an *absorbing* state: a state which, once entered, is never exited – i.e. $\rho_{ii} = 1$. One might also consider locally- (or semi-) absorbing states, whereby the process resides in one state for a 'sufficiently' long time but not necessarily with complete probability degeneracy (i.e., $\rho_{ii} \approx 1$) – this is perhaps closer to a conventional time-series break definition. This definition of structural break as a (semi) permanent change of state is also related to the existence of Hysteresis defined as a system with multiple equilibrium. Once unemployment suffers a rapid increase or decrease, it tends to stay in the new state (lock-in). This is probably a closer definition of the Hysteresis arising in theoretical models of labour markets in transition economies.

Examining the smoothed probabilities for each country (Figure 2), most countries have indeed spent long periods in one state.[27] The Czech Republic spent the early sample (up until around 1996) in the first (low-unemployment) state followed by a transition to a high-unemployment one. We might therefore tentatively suggest a break around 1997–98 (as Table 4 suggests). As commented earlier on, this may be due to the stabilization plan introduced in 1997. The same can be said for Lithuania with a likely break in 1998–1999 (although Table 4 picks up the earlier break of 1997:1).[28] Slovakia appears to have spent most of its time in the (low-unemployment) first state but from 1998 onwards appears to head permanently into a high unemployment state (Table 4 tends to pick up the break around late 1992). These possible two breaks are consistent with the more hesitant recovery of Slovakia in the mid 1990s that can be observed in Figure 1. Latvia appears to have entered a high-unemployment state in the immediate aftermath of the Russian crisis (also indicated by Table 4) but starts to have recovered by around mid 2000, in line with the recovery of the Russian economy as well.[29]

These results hence show that, for several transition countries, unemployment appears to follow a multiple equilibrium pattern: shocks that affected unemployment during the last years of the past decade seem to have moved these economies towards a high-unemployment equilibrium. Hysteresis, although not manifested in general as a linear random-walk process behaviour, seems to take the form of multiple equilibria, especially for countries such as the Czech Republic, Lithuania and Slovakia. This lock-in pattern behaviour is supportive of recent models of transition in labour markets such as Boeri (2001).

[27] The comparison between absorbing states and time-series structural breaks is by no means exact. The former essentially verify a break when there is complete degeneracy (i.e. there is no further exit from state) whilst the latter may be more commonly thought to register a structural break during the transition away from a previous state; that is to say, as $\rho_{ii} \rightarrow 1$. Notwithstanding, comparing time-series (as earlier examined) and Markov-switching 'break' detection methods may be a useful cross-checking exercise.

[28] Lithuania represents the only case of a fully absorbing state since $\rho_{22} = 1$.

[29] Some caution, however, should be borne in mind concerning breaks and regime shifts towards the end of the sample. In our cases, most of the breaks correspond well across methodologies and appear to match well, historical events – for example, large well-dissipated shocks like the Russian crisis. Nevertheless, it is of course only feasible to verify breaks and regime shifts using very much longer samples. Thus, our discussion is necessarily tentative.

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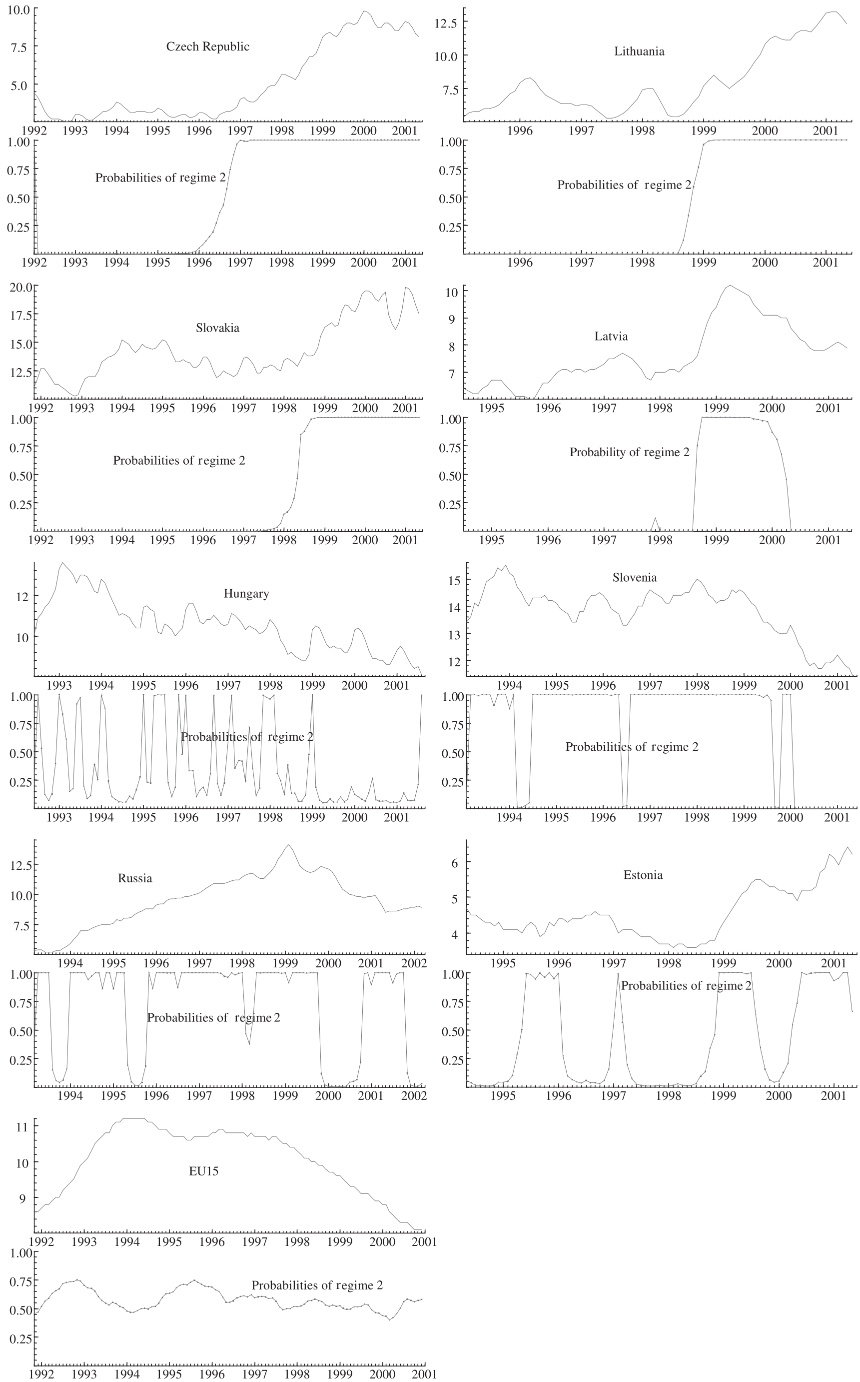


Figure 2. Unemployment rates and smoothed probabilities.

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## V Conclusions

In this paper, we have undertaken a systematic analysis of the dynamic behaviour of unemployment in transition economies benchmarked against the EU-15. We tested for the existence of hysteretic features in labour markets making use of both unit-root tests and Markov-switching regressions. Our findings show that, in general, we can reject the unit-root hypothesis after controlling for structural changes and business-cycle effects, but that we can observe the presence of a high and low unemployment equilibria towards which the economy fluctuates after sufficiently large shocks.

When compared with the behaviour of unemployment dynamics in the EU during the past decade, we can see that transition countries' unemployment shows a faster speed of adjustment and larger changes in unemployment equilibria across regimes. Exception to this pattern would be Croatia, whose unemployment behaviour is best described as a linear unit root or near unit-root process, and Latvia, where unemployment seems to follow a random walk and regime changes. For the other countries, unemployment persistence is relatively low, which is consistent with less regulated labour markets. Moreover, for several countries we find that once unemployment shifts towards a new regime, it tends to remain locked into it – or, at least, remain there for a long period. Notable cases of this during the final years of the 1990s are the Czech Republic, Lithuania and Slovakia. We can thus conclude that unemployment dynamics in Eastern Europe are characterized by a switching unemployment equilibrium towards which actual unemployment reverts quicker than in the EU. This pattern is supportive of recent theoretical models of the labour market in transition countries.

These results have important implications for labour market reforms, as well as macro-stabilization policy in the CEECs. Standard progressive macroeconomic stabilization policies do not appear to have a long-lasting impact on unemployment, at least not longer than what the EU experience reveals. However, deeper reforms of both labour and goods markets, (which might constitute 'large' shocks) that are likely to continue in the CEECs should take into account the possibility of having a long-lasting impact on the equilibrium level of unemployment.

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## Appendix: Data Sources

Country, data sources and samples are:

**Poland:** Central Statistical Office of Poland [Jan. 1991–June 2001]
**Romania:** National Commission for Statistics [Dec. 1991–Apr. 2001]
**Slovenia:** Central Bank of Slovenia [Jan. 1992–May 2001]
**Croatia:** Statistical Office of Croatia [Jan. 1992–May 2001]
**Hungary:** Central Statistical Office of Hungary [May 1991–Aug. 2001]
**Bulgaria:** WIIW, Eastern Europe Economy [Jan. 1991–June 2001]
**Czech Republic:** WIIW, Eastern Europe Economy [Jan. 1991–May 2001]
**Slovak Republic:** Slovak Statistical Office [Jan. 1991–May 2001]
**Estonia:** OECD Main Economic Indicators [May 1993–May 2001]
**Latvia:** Latvijas Statistiskas/Monthly Bulletin [Jan. 1994–May 2001]
**Lithuania:** Lithuanian Department of Statistics [Jan. 1994–May 2001]
**Russia:** Goskomstat/Russian Economic Trends [Jan. 1992–Mar. 2002]
**EU-15:** EUROSTAT [Jan. 1991–Dec. 2000]
**Note:** WIIW = *Wiener Institut für Internationalen Wirtschaftsvergleich*

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